Farm-Level Production Effects from Participation in Government Commodity Programs: Did the 1996 Federal Agricultural Improvement and Reform Act Make a Difference?

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In theory, lump-sum transfers are a way to redistribute wealth without distorting production decisions. Recent agricultural policy reforms are an unusual large-scale application of this concept. The 1996 Federal Agricultural Improvement and Reform (FAIR) Act removed most price-contingent agricultural subsidies and replaced them with Production Flexibility Contracts—lump-sum payments with few ties to farmers' production decisions. The payments were envisioned as way to maintain income transfers to agricultural interests while minimizing production distortions. The United States has argued that these "decoupled" agricultural payments are minimally trade distorting (USDA-ERS).

There is a great deal of uncertainty about how the system of agricultural payments established under the 1996 reforms affects production. Some, including representatives from developing nations with domestic agricultural sectors that compete with the United States, contend that decoupled payments significantly affect agricultural production and trade. Two general arguments underpin this assertion. First, decoupled payments are not really lump sum—that is, despite the reforms, important links to production remain. For example, there are restrictions in the FAIR Act that preclude new vegetable plantings or the conversion of land to nonagricultural uses. Second, the perfect market assumptions underpinning the theory that lump-sum payments do not affect production are not maintained in practice. The existence of labor, credit, and insurance market imperfections—departures from the simple neoclassical model—may give rise to production effects (USDA-ERS).

Empirical estimates of production distortions resulting from decoupled payments are usually based on theoretical assumptions about how producer attitudes toward risk changes with wealth (e.g., Chavas and Holt; Hennessy; Mullen et al.; Sumner; Young and Westcott). An exception is a study by Goodwin and Mishra that estimates production effects from decoupled payments using cross-sectional data from the Heartland region over 1998–2001. Others argue that direct payments may help credit-constrained farmers to remain in production (Chau and de Gorter; Roe, Somwaru, and Diao). Nevertheless, empirical evidence on the relationships between wealth, risk, and agricultural production is scarce. In addition, the utility theory that underlies the estimated production effects from decoupled payments is often contradicted in the experimental literature (e.g., Arrow et al.; Kahneman and Tversky). Wealth-related production effects stemming from decoupled income support that have been considered in the theoretical literature thus far are likely to be small. Larger production distortions may arise if liquidity constraints are large and pervasive or if farmers make decisions based not on strict adherence to the incentives embodied by current policy parameters, but on what they expect future policies may be, or what has succeeded for them in the past. These possibilities need further study.

Obtaining direct observational evidence of production effects from government programs is a challenge, chiefly because there is no

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obvious control group against which to compare the performance of program participants. The federal government has been subsidizing key agricultural commodities since the Great Depression and farmers have faced the same eligibility requirements for program participation. As a result, virtually all farmers growing these key agricultural commodities have been affected by government programs, either directly or indirectly. In addition, program rules have changed markedly from one farm Act to the next, but so have commodity prices, crop yields, technologies, and many other factors, so comparisons of aggregate data over time are not straightforward. Indeed, changes in program rules have arguably been influenced by changes in prices. These reasons, among others, create difficulties for measuring the effect of federal agricultural program payments on U.S. production of commodity crops before and after the 1996 Farm Act.

In this article, we estimate at a microeconomic level the production effects of federal commodity programs using a unique farm-level panel dataset derived from microfiles of the 1987, 1992, 1997, and 2002 U.S. Agricultural Censuses. These data include information on the amount of land allocated to particular crops, total government payments, and land set aside in accordance to program requirements. From these datasets, we identify a substantial number of farms that have not participated in government programs. Our empirical approach is to compare growth in program crop acreage between consecutive Censuses of farms participating in government programs with similar farms that were not participating—that is, we use nonparticipating farms as a control group. We also compare the changes in farm size and farm sales of participants and nonparticipants. Separate comparisons are made during periods of coupled payments (1987–1992), decoupling (1992–1997), and (mostly) decoupled payments (1997–2002). If the 1996 FAIR Act effectively removed incentives for farmers to overproduce and the nonparticipant group is a viable control, we would expect the effect of participation on program acreage in the 1987–1992 coupled period to be larger than in the 1997–2002, comparatively decoupled period.

The crux of this analysis is the assumption that nonparticipants are a viable control group. Since farms are not randomly assigned to participate in farm programs, the main empirical challenge is to control for unobserved factors that may have influenced both program participation and plantings of program crops. Nevertheless, similar farms may have made different participation decisions due to heterogeneous expectations about future prices and government programs, and perhaps heterogeneous costs associated with compliance provisions initiated in 1985.[1] A historical decision not to participate may have had persistent effects on the profitability of participation, because program benefits were linked to historical participation.[2] Thus, it is plausible that conditioning on observable factors, participating and nonparticipating farms are otherwise similar.

We control for time-invariant unobserved heterogeneity of farms by analyzing farm-specific changes in program-crop acreage between Census periods. We control for factors other than participation that could affect growth rates using a large set of fixed effects associated with farm type, scale, location, and operator age. These controls enable comparisons of program participants and nonparticipants that are observationally similar.

Data and Summary Statistics

Data on farm and operator characteristics are from the farm-level files of the 1987, 1992, 1997, and 2002 Agricultural Censuses maintained by the National Agricultural Statistics Service (NASS) of the U.S. Department of Agriculture. Conducted every five years, the Agricultural Census includes essentially all U.S. farms.

Our sample is restricted to operations defined as farming in two consecutive Census years. We thus obtain three panels: 1987–1992, 1992–1997, and 1997–2002. We restrict each panel to farms with at least 10 acres harvested of program crops in the first period of the panel, with "program crops" including corn, wheat, barley, oats, cotton, rice, and sorghum. These restrictions reduce the sample

[1] Since 1985, farmers with highly erodible land have been required to use soil-conserving practices in order to retain eligibility for farm programs. See Claassen et al.

[2] Prior to the 1996 Act, participation in government programs constrained farmers' production decisions in a number of ways. Planted acreage of program crops was limited to a historical "base" of acreage, equal to a five-year rolling average of historical plantings. Under acreage reduction programs, farmers were required to set aside (leave fallow) a share of their base. This share varied from year to year. Various programs allowed farmers to maintain base and a share of their payments if they elected to set aside a larger share of their base than the minimum amount required.

to 453,948, 381,794, and 313,679 observations in the three respective panels.

Because of policy changes brought by the 1996 FAIR Act, we use two indicators of program participation: prior to 1996, participation is indicated by positive set-aside acres; after 1996, participation is indicated by receipt of government payments. Prior to the Act, payments were linked to commodity prices, so many participating farms did not receive payments, and the existence of set-aside acreage is a clearer indicator of participation.[3] Since the 1996 FAIR Act, there have been no set-asides, but all participating farms have received Production Flexibility Contracts, the (nearly) lump-sum transfers described in the introduction.[4] We thus use receipt of government payments as an indicator of participation in 1997 and 2002.[5] To clearly delineate between participants and nonparticipants, we drop all farms that participated in one year of each panel and did not participate in the other year.[6]

With these restrictions, the three panel datasets consist of 314,724, 251,287, and 216,521 observations, of which 37%, 33%, and 22% were nonparticipants in the 1987–1992, 1992–1997, and 1997–2002 panels, respectively (table 1). While the overall number of farms declines across the three panels, the number of nonparticipants declined by more than the number of participants. The number of nonparticipating farms in each size class declines over each consecutive five-year period and drops by 59% overall from the first to the last panels (1987–2002). This compares to an overall 15% decline in the number of participants during the same period.

For each panel, participants and nonparticipants were categorized into five groups based on program-crop acreage. The acre cutoffs for the five groups were determined by the 40, 60, 80, and 90% quantiles from the initial year's distribution. We chose these cutoff points, rather than standard quintiles (20, 40, 60, and 80%), because the distribution of farms is highly skewed toward larger farms.[7] Table 1 shows that the distribution of nonparticipating farms is highly skewed toward smaller acreage quantiles, while participating farms are more evenly distributed across size classes. For example, the 1987–1992 panel includes 83,593, 18,250, 10,046, 3,006, and 2,301 nonparticipants in the first to fifth program acreage quantiles, respectively. In comparison, there were 42,702, 40,675, 54,092, 29,811, and 30,248 participants in these same categories.

Table 2 compares the average values of key variables for participants and nonparticipants of different sizes in the three panels. While nonparticipants had less program-crop acres to begin with compared to participants, their program-crop acreage declined by more over time in percentage terms.[8] Over 1987–1992, depending on the acreage category, nonparticipants declined in program-crop acreage by 37% to 61% while participants increased in program-crop acreage by 4% to 24% on average. Over the two following periods, both participants and nonparticipants decreased in program-crop acreage but nonparticipants' acreage declined by substantially more. Depending on the size category, nonparticipants' program-crop acreage decreased by 43–74% and 63–87% while participants declined by only 12–18% and 15–33% over 1992–1997 and 1997–2002, respectively.

The summary statistics in tables 1 and 2 suggest that participation in government programs could have had large effects on farm-size growth and farm-level plantings of program-related crops, both before and after the 1996 FAIR Act. Nevertheless, participants and nonparticipants differ somewhat in their characteristics: nonparticipating farmers tend to be older and their operations are smaller, more specialized in cattle, and less specialized in corn production, even after conditioning on program-crop acreage (table 2). The comparisons, even within program-crop acre classes, therefore may be indicative of differences in farm and operator types that are associated with participation, rather than participation itself. We begin to explore this possibility using multiple regression analysis.

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[3] Some payments remained linked to prices and production after 1996; however, set-asides were not used to allocate these payments.

[4] Some farms in some years also received marketing loans or ad hoc market loss assistance.

[5] Government payments are measured as total direct government payments reported in the Census minus payments under the Conservation and Wetland Reserve programs, which are separately specified.

[6] This measure of nonparticipation also excludes farmers who did not participate in government commodity programs, but received some other direct government payments in the form of disaster assistance.

[7] This breakdown ignores failed acres (planted but not harvested) as the Census does not report failed acres by crop. Because the quantile cutoffs were calculated from all farms in the Census, the distribution across quantiles of the farms remaining in our sample differs slightly from 20, 40, 60, and 80%.

[8] As described in the next section, change in program crop acres is scaled from –200% to 200% so as to equate the absolute value of percentage increases and decreases.

Table 1. Distribution of Farms

Classification, by Period and Program Crop Acreage	Participants (No., % in Each Class)	Nonparticipants (No., % in Each Class)	Total Farms (No., % in Each Class)	Participants as % of All Farms in Acreage Class	Nonparticipants as % of All Farms in Acreage Class
1. 1987–1992 panel (>=10 program acres in 1987)	197,528 (100)	117,196 (100)	314,724 (100)	62.8	37.2
a. 10–70 program acres in 1987	42,702 (21.6)	83,593 (71.3)	126,295 (40.1)	33.8	66.2
b. 71–136 program acres in 1987	40,675 (20.6)	18,250 (15.6)	58,925 (18.7)	69.0	31.0
c. 137–285 program acres in 1987	54,092 (27.4)	10,046 (8.6)	64,138 (20.4)	84.3	15.7
d. 286–480 program acres in 1987	29,811 (15.1)	3,006 (2.6)	32,817 (10.4)	90.8	9.2
e. >480 program acres in 1987	30,248 (15.3)	2,301 (2.0)	32,549 (10.3)	92.9	7.1
2. 1992–1997 panel (>=10 program acres in 1992)	166,670 (100)	84,617 (100)	251,287 (100)	66.3	33.7
a. 10–88 Program acres in 1992	35,414 (21.2)	60,793 (71.8)	96,207 (38.3)	36.8	63.2
b. 89–183 program acres in 1992	35,589 (21.4)	12,478 (14.7)	48,067 (19.1)	74.0	26.0
c. 184–395 program acres in 1992	45,731 (27.4)	6,871 (8.1)	52,602 (20.9)	86.9	13.1
d. 396–650 program acres in 1992	25,089 (15.1)	2,441 (2.9)	27,530 (11)	91.1	8.9
e. >650 program acres in 1992	24,847 (14.9)	2,034 (2.4)	26,881 (10.7)	92.4	7.6
3. 1997–2002 panel (>=10 program acres in 1997)	168,350 (100)	48,171 (100)	216,521 (100)	77.8	22.2
a. 10–99 program acres in 1997	45,703 (27.1)	36,253 (75.3)	81,956 (37.9)	55.8	44.2
b. 100–200 program acres in 1997	36,534 (21.7)	6,120 (12.7)	42,654 (19.7)	85.7	14.3
c. 201–450 program acres in 1997	42,360 (25.2)	3,371 (7.0)	45,731 (21.1)	92.6	7.4
d. 451–760 program acres in 1997	21,506 (12.8)	1,275 (2.6)	22,781 (10.5)	94.4	5.6
e. >760 program acres in 1997	22,247 (13.2)	1,152 (2.4)	23,399 (10.8)	95.1	4.9

Notes: Program-crop acres equal the sum of barley corn, cotton, oats, rice, sorghum, and wheat acres harvested. "Participants" had set-aside acreage or received direct government payments in both consecutive Census years. "Nonparticipants" had no set-aside acreage and no direct government payments in both years. Those participating in one year and not the other are excluded from this analysis. The acres cutoffs for the five groupings for each panel were determined from the 40, 60, 80, and 90% quantiles from the initial year's distribution of farms with 10 or more acres harvested in program crops. Because the quantile cutoffs were calculated from all farms in the Census, the distribution across quantiles of the farms remaining in our sample differs slightly from 20, 40, 60, and 80%.

Table 2. Summary Statistics by Quantile and Participation Category

	1987–1992		1992–1997		1997–2002	
Panel Years	Nonparticipants	Participants	Nonparticipants	Participants	Nonparticipants	Participants
1. First quantile of program acres in initial year						
%Δ Program acres	−61.01	24.83	−73.74	−18.14	−86.52	−32.38
Beginning-year program acres	31.91	42.10	35.63	50.73	36.80	52.21
Beginning-year corn acres	17.49	29.70	20.53	34.79	22.23	35.55
Beginning-year total expenditures ($)	71,383.36	71,069.79	105,774.90	106,480.70	129,213.12	144,423.96
Beginning-year cattle (head)	66.42	54.71	67.35	71.33	60.89	89.44
Operator age in ending year	55.62	53.57	56.25	56.29	57.03	56.72
2. Second quantile program acres in initial year						
%Δ Program acres	−43.28	20.31	−61.10	−13.78	−72.45	−18.30
Beginning-year program acres	97.23	101.94	124.64	132.95	139.48	146.00
Beginning-year corn acres	52.15	72.21	69.04	93.27	78.76	102.30
Beginning-year total expenditures ($)	124,042.44	107,940.33	176,396.66	141,072.17	251,563.82	218,287.45
Beginning-year cattle (head)	122.61	82.39	125.27	91.03	127.51	119.16
Operator age in ending year	54.99	52.02	55.51	54.37	56.37	55.76
3. Third quantile program acres in initial year						
%Δ Program acres	−40.20	17.05	−52.62	−12.47	−62.78	−14.91
Beginning-year program acres	189.82	200.38	260.90	274.67	304.14	310.15
Beginning-year corn acres	91.18	131.94	127.52	182.63	156.78	204.40
Beginning-year total expenditures ($)	201,755.71	137,612.92	308,480.30	194,015.43	519,868.15	282,947.50
Beginning-year cattle (head)	205.25	106.91	228.43	120.86	307.80	148.70
Operator age in ending year	54.81	51.38	54.93	53.11	55.34	54.78
4. Fourth quantile program acres in initial year						
%Δ Program acres	−37.29	11.89	−51.40	−12.69	−60.19	−15.34
Beginning-year program acres	361.32	367.76	497.25	505.09	578.35	586.46
Beginning-year corn acres	145.01	195.50	197.58	281.77	258.05	325.15
Beginning-year total expenditures ($)	366,304.10	186,406.36	307,434.20	234,580.51	474,437.47	349,195.80
Beginning-year cattle (head)	387.82	156.56	296.01	162.29	311.25	190.24
Operator age in ending year	55.05	51.19	54.17	52.54	54.50	54.23
5. Fifth quantile program acres in initial year						
%Δ Program acres	−40.48	3.59	−43.77	−12.50	−64.75	−20.73
Beginning-year program acres	924.41	876.15	1,264.52	1,150.67	1,493.71	1,399.55
Beginning-year corn acres	222.92	249.29	316.00	364.34	424.43	448.88
Beginning-year total expenditures ($)	507,036.45	267,029.19	614,002.80	361,832.26	775,101.51	527,288.36
Beginning-year cattle (head)	469.11	266.39	379.52	275.47	485.11	335.91
Operator age in ending year	54.42	51.68	54.16	52.62	54.91	54.17

Multiple Regression Estimates

Our goal is to predict three indicators of farm performance: growth in program crops, growth in farm size (total land area), and growth in sales, using a participation indicator and controls for farm types and locations.[9] By examining growth rates rather than levels (i.e., first differencing), we remove all time-invariant heterogeneity across farms. All three growth measures are measured in percent changes, calculated as $\Delta Y_i = 200$ (*value in period 2 – value in period 1*)/(*value in period 1* + *value in period 2*). Calculating the percentage changes in this way makes the absolute value of increases equal to decreases. It is also bounded between –200% and 200%, which attenuates the influence of outliers. For example, a change from 1 acre to 1,000 acres is calculated as a change of 199.6% and a change from 1,000 to 1 acre is calculated as a change of –199.6%.

Each regression estimates a model of the form,

$$\text{(1)} \qquad \Delta Y_i = \delta P_i + \mathbf{X}_i \boldsymbol{\beta} + \varepsilon_i,$$

where P_i is a participation indicator variable, $\mathbf{X}_i$ is a vector of control variables, and $\boldsymbol{\beta}$ and δ are parameters to be estimated. To allow for heterogeneity in growth rates, we control for farm and operator characters that are likely to be associated with farm profitability or life-cycle investment decisions, such as farm location, farm type, farm size, and operator age (Kimhi and Bollman; Sumner and Leiby; Weiss; Zepeda). In particular, we consider the following controls:

State : Fixed effects (dummy variables) for each U.S. state;
SIC class : Fixed effects for each of twenty-one 6-digit Standard Industrial Classification (SIC) codes;
Program-acre class : Fixed effects for each of the five quantiles of total program acres, as delineated in table 1;
Age class : Fixed effects for each of five age-class quantiles, delineated from the 20, 40, 60, and 80% of the age distribution of farm operators in the sample;
Size class : Fixed effects for each of five farm-size classes, where size is defined by total land in farm, delineated from the 40, 60, 80, and 90% of the size distribution of farms;
Soybean acres : Soybean acres harvested in the beginning year;
Corn acres : Corn acres harvested in the beginning year;
Irrigated acres : Irrigated acres in the beginning year;
Cattle: Head of cattle plus calves in beginning year, including head sold plus inventory;
Hogs : Hogs in beginning year, including head sold plus inventory.

We consider the following sets of interactions of these control factors:

1. Program-Acre Class × Age Class × Size Class × SIC Class
2. Program-Acre Class × State × Size Class
3. Soybean Acres × Program-Acre Class × Age Class × Size Class
4. Corn Acres × Program-Acre Class × Age Class × Size Class
5. Irrigated Acres × Program-Acre Class × Age Class × Size Class
6. Cattle × Program-Acre Class × Age Class × Size Class
7. Hogs × Program-Acre Class × Age Class × Size Class

The categorical variables narrow the source of identification to differences between participants and nonparticipants within each group. By interacting sets of fixed effects variables, we parse farms into smaller and more homogenous groups. For example, the first set of interactions includes a dummy variable for each program-acre class, age class, size class, and SIC class combination. This narrows the source of identification to farms within the same SIC code, same size class, same age class, and same program class, and includes a total of $5 \times 5 \times 5 \times 21 = 2{,}625$ dummy variables.[10] Interacting the continuous variables with the classification

[9] Researchers have debated the merits of different measures of farm performance (Mishra and Morehart). In addition to the indicators examined here, possible measures include net farm income, net farm income per dollar of assets, and total returns from farming (Mishra et al.).

[10] When we consider all sets of interactions simultaneously, some redundancy among the dummy variables may cause the degrees of freedom to be somewhat less than this calculation.

Table 3. Multiple Regression Estimates and Analysis of Variance

		Estimated Effect (DF)	Regression Sum of Squares
1. 1987–1992 Panel			
Estimated effects of participation			
a. %Δ program acres	$R^2 = 0.255$	59.1 (1)	14,539.94
b. %Δ land in farms	$R^2 = 0.093$	14.6 (1)	889.41
c. %Δ sales	$R^2 = 0.137$	21.6 (1)	1,933.74
Controls & analysis of variance for (a)%Δ program acres			
d. Program-acre class × age class × size class × SIC class		(2,703)	6,311.10
e. Program-acre class × state × size class		(1,029)	4,433.68
f. 1987 soybean acres × program-acre class × age class × size class		(125)	1,240.57
g. 1987 corn acres × program-acre class × age class × size class		(125)	153.64
h. 1987 irrigated acres × program-acre class × age class × size class		(125)	283.01
i. 1987 cattle × program-acre class × age class × size class		(125)	150.61
j. 1987 hogs × program-acre class × age class × size class		(125)	103.16
2. 1992–1997 panel			
Estimated effects of participation			
a. %Δ program acres	$R^2 = 0.181$	40.0 (1)	5,391.23
b. %Δ land in farms	$R^2 = 0.077$	14.3 (1)	688.27
c. %Δ sales	$R^2 = 0.127$	24.1 (1)	1,943.36
Controls & analysis of variance for (a)%Δ program acres			
d. Program-acre class × age class × size class × SIC class		(2,577)	3,556.13
e. Program-acre class × state × size class		(973)	5,646.71
f. 1992 soybean acres × program-acre class × age class × size class		(125)	1,040.34
g. 1992 corn acres × program-acre class × age class × size class		(125)	164.14
h. 1992 irrigated acres × program-acre class × age class × size class		(125)	225.78
j. 1992 cattle × program-acre class × age class × size class		(125)	131.63
k. 1992 hogs × program-acre class × age class × size class		(125)	100.41
3. 1997–2002 panel			
Estimated effects of participation			
a. %Δ program acres	$R^2 = 0.222$	38.0 (1)	3,742.64
b. %Δ land in farms	$R^2 = 0.084$	15.8 (1)	645.03
c. %Δ sales	$R^2 = 0.106$	22.5 (1)	1,311.36
Controls & analysis of variance for (a) %Δ program acres			
d. Program-acre class × age class × size class × SIC class		(2,485)	6,400.76
e. Program-acre class × state × size class		(978)	4,568.89
f. 1997 soybean acres × program-acre class × age class × size class		(125)	1,336.07
g. 1997 corn acres × program-acre class × age class × size class		(125)	162.38
h. 1997 irrigated acres × program-acre class × age class × size class		(125)	125.81
j. 1997 cattle × program-acre class × age class × size class		(125)	528.77
k. 1997 hogs × program-acre class × age class × size class		(125)	93.88

Notes: Standard errors for the effect of participation are less than one quarter of 1% in all cases, as implied by the *F*-tests. Due to space limitations, the coefficients for the many fixed effects are not reported. The Regression Sum of Squares reports the increase in fit obtained by the associated set of factors relative to the full model. The reported *F*-statistics are for the joint significance of the associated set of factors. All sets of factors, except for 2-k and 3-k are significant with a p-value less than 0.001. *F*-tests for the controls in the %D. "Land in farms" and "% Δ sales" regressions are not reported due to space limitations.

variables (in 4, 5, 6, and 7), allows these variables to have different marginal effects within each subgroup delineated by the class variables with which they are interacted. These seven sets of control variables account for a wide range of heterogeneity in farm size and program-acre growth.

Separate regressions are estimated for each of the three two-year panels, 1987–1992, 1992–1997, and 1997–2002. Table 3 reports the estimated effect of participation on program-acre growth, farm-size growth (measured in land area), and sales growth, the R^2 for each regression, *F*-tests indicating the statistical significance of the participation indicator in each regression, and *F*-tests for the significance of each set of control factors in the three program-acre growth regressions. *F*-tests for the control factors in the farm-size growth and sales growth regressions are omitted for brevity.

In all regressions, participation has a large positive estimated effect on growth. The estimated effect on the change in program acres is largest (59.1 percentage points) in the early panel (1987–92). Much of this effect likely stems from a large reduction in set-aside acreage. In 1987, mandatory set-asides were 35%, 27.5%, and 25% of program acreage for rice, cotton, and wheat, respectively, and 20% for corn, sorghum, barley, and oats. In 1992, these amounts were reduced to 5% for all crops except for rice (10%), and cotton and oats (0%). If approximately 20 percentage points of the difference between participants and nonparticipants is explained by changes in set-asides, the remainder is approximately equal to the estimated effects in the 1992–97 and 1997–2002 panels, 40% and 38%, respectively. Estimates for the effect of participation on farm-size growth and sales growth are also similar across time, equal to 14.6, 14.3, and 15.8 percentage points for size growth, and 21.6, 24.1, and 22.5 percentage points for sales growth. All of these estimates are stable across a wide range of alternative specifications of the control variables.[11]

In all regression models, participation is by far the most important factor predicting all three measures of farm performance. For program-acre growth, this is illustrated by the *Regression Sum of Squares*. In panels (1) and (3), the participation indicator explains a larger share of the variance in program-acre growth than the approximately 2,500–2,700 variables embodied in the first set of control variables; in panel (2), the participation indicator explains almost as much of the variance as these variables.

Although the participation indicator is not associated with obvious unobservable indicators of farm type or location, the possibility of sample selection bias can be addressed using a treatment effects model that allows for possible correlation between the errors of separate program participation and growth equations (Greene, p. 713). Preliminary results from this line of analysis, which are available from the authors upon request, indicate the effects of program participation on farm growth that are similar in magnitude to the effects reported table 3.

[11] Limited space prevents us from reporting these results here. A full set of estimates is available from the authors upon request.

Conclusion

This article studies the effects of large-scale U.S. federal agricultural programs by comparing program participants to nonparticipants that are otherwise similar in their observed characteristics. The results indicate that participants increased plantings of program crops considerably (38–59 percentage points) more than nonparticipants. The relative increase is about the same size both before and after the 1996 FAIR Act, after accounting for the reductions in participants' set-aside acreage. Estimated effects of participation on farm-size growth and total sales growth were also consistent across time, increasing 14–16 percentage points and 22–24 percentage points, respectively. If the observed associations are causal, these findings suggest that participation in domestic agricultural programs, including the mostly decoupled 1997 payment regime, had significant farm-level production effects. Although the study is broader than the issue of decoupling, the results suggest decoupled payments affect production at the farm level.

The estimates from this study do not indicate the aggregate magnitude of the distortion caused by domestic agricultural payments, even if the identified group of nonparticipants provides a viable control group for measuring program impacts. A marked decline in plantings of program-related crops would likely cause prices to rise, attenuating the overall decline in aggregate plantings. In the absence of agricultural program payments, we thus expect acreage would change by a smaller amount than our farm-level estimates for the effect of program participation. Moreover, our results should be viewed as preliminary. To our knowledge, this is the first attempt to examine effects of federal agricultural programs by comparing participants with nonparticipants. More work is needed to verify an absence of selection bias and determine more precisely the mechanisms through which these programs may be influencing production decisions.

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